Gender Bias and Quantity Quality Tradeoff of Children in China

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A Thesis Submitted in Partial Fulfilment
of the Requirement for the Degree of
Master of Philosophy
in
Economics

Abstract

This thesis uses education as a dependent variable to examine the intra-household distribution of resources to children in China from the following two aspects: (1) the gender bias, (2) the quantity-quality tradeoff of children. We find that gender bias against girls only exists in rural China and only for children of the older birth cohorts. We also find that for the rural sample, female-headed households tend to favor the education of girls, and both the education level of fathers and higher household income per capita help to reduce the gender bias. However, we do not find that having the second or more children or the birth order of children changes the gender bias. Finally, we examine whether there is a quantity-quality tradeoff of children. By directly employing the one-child policy variable as an instrumental variable, we find that there is indeed a quantity-quality tradeoff of children.

摘要

本文使用教育作爲因變量，從以下兩方面分析中國家庭如何將內部資源分配於其子女身上：（1）性別歧視；（2）孩子的數量和素質之間的權衡。我們發現，對女孩子的性別歧視僅存在於中國農村，並且是農村中較早出生的那群孩子中。同時，我們發現在農村的樣本中，女性戶主傾向于讓女生接受更多的教育，而且，父親的教育水平及家庭的人均收入越高，性別歧視也會越少。然而，我們並不能證明二胎或者更多的孩子，以及孩子的出生次序會改變性別歧視。最後，我們檢驗是否存在孩子的數量和素質之間的權衡。通過引入獨生子女政策作爲工具變數，我們發現確實存在這種孩子數量和素質的權衡。

Acknowledgement

I would like to thank my supervisor, Prof. Li Hongbin, for his valuable guidance and support. Without his patient and helpful supervision, I could not have finished this thesis. I would also like to thank Prof. Zhang Junsen and Prof. Leung Ka Yui for their constructive comments and suggestions which greatly improve the quality of my thesis. Thanks to my fellow classmates, friends and family members especially Cheung Wai Ki, Huang Ying, Meng Lingsheng, Moy Hon Man, Tang Lok Ming, Wong Chung Yiu, Wong Loong Kuen and Xue Jianpo, who have given me infinite help and encouragement during this stressful period of doing thesis research. Thanks to the Carolina Population Center for kindly providing the data

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1 Introduction

Parents' decisions on intra-household allocation of resources among children have a great impact on child development, health, education, which will in turn affect the future income of children (Haveman and Wolfe, 1995; Danziger and Waldfogel, 2000; Young, 2002). With uneven intra-household distribution, higher household income does not necessarily mean that all children are equally better off (Haddad and Kanbur, 1990; Browning and Lechene, 2001; Quisumbing, 2003). In this thesis, we investigate intra-household allocation in China using child educational attainment as an indicator. In particular, we examine whether there is a gender bias in China in terms of the years of education children receive, how household characteristics and child birth order affect the magnitude of the gender bias, and whether the number of children affects the educational attainment of each child.

How do parents allocate household resources among children greatly depends on the value of each child to parents. The value varies with a child's gender and age, the parents' education levels, as well as other child and household characteristics (Ahn, 1995). When gender is a consideration, parents may very likely allocate more resources to children of one gender, normally boys, than the other gender. We call this gender bias. In China, boys are very likely to be preferred by parents and thus receive more education because boys are responsible for caring aging parents and because boys can carry on the family name. Moreover, as women's income is widely believed to be significantly lower than men's, investing in boys may have a higher return even if both boys and girls share equally the duty of caring aging parents.

There is a huge empirical literature that attempts to measure intra-household gender bias among children in developing countries, but no consensus has been reached so far. Some find that the gender bias exists in dimensions such as nutrition, health, and education (see e.g., Drèze and Sen, 1989; Harriss, 1990; Dasgupta, 1993; Strauss and Thomas, 1995; Ilahi, 2001; Kingdon, 2004). Others (eg. Subramaniam (1996) and Deaton (1998)) do not detect gender bias.

Our first objective in this thesis is to examine whether there is a gender bias in China in the late 1980s and early 1990s. Drawing on China Heath and Nutrition Survey data collected by the Carolina Population Center, we find that the gender bias in education attainment against girls only existed for the 1989 sample, but not for the 1991 and 1993 samples. Moreover, we only detect gender bias for the older birth cohorts of the 1989 sample. For younger cohorts, girls even had more education than boys. Further analysis using the urban and rural sample separately shows that all the detectable gender bias is attributable to the rural sample.

In addition to examining gender bias across age groups, we are also interested in whether family characteristics and child birth order, which may affect resource allocation on child education, have significant effects on gender bias. We find that for the rural sample, female-headed households tend to favor the education of girls, and both the education level of fathers and higher household income per capita help to reduce the gender bias. However, we do not find that having one more child or the mother's education affects the magnitude of gender bias. Our further analysis shows that although both the first and last births in a family receive more education than their siblings, girls of the first and last births are neither favored nor discriminated in terms of education.

The second issue we investigate in this thesis is the quantity and quality tradeoff of children. The quantity-quality model, first introduced by Becker (1960) in 1960, suggests that the marginal cost of child quality rises with family size. One prediction of the model is that the increase of the total number of children will reduce the average quality of all children. However, as education and the number of children are both endogenous variables chosen by parents, the main difficulty of investigating the tradeoff is to identify the true casual effect of quantity on quality. For example, if families planned to invest less in child education also preferred to have more children, the negative correlation between quantity and quality estimated by ordinary least square regressions would not be their true casual relationship. In this thesis, we use the birth control policy as an instrumental variable to solve this endogeneity problem.

We indeed find that a quantity and quality tradeoff of children.

The structure of this thesis is organized as follows: Section 2 describes the data set and the main variables. Section 3 investigates the gender and birth-order effects on individual child's education, while section 4 analyzes the quantity-quality tradeoff. Section 5 concludes.

2 Data

In this thesis, we use the China Health and Nutrition Survey (CHNS) data, which were collected by the Carolina Population Center (CPC) at the University of North Carolina at Chapel Hill, the Institute of Nutrition and Food Hygiene, and the Chinese Academy of Preventive Medicine. The sample households were randomly drawn from eight provinces including Liaoning, Shandong, Jiangsu, Henan, Hubei, Hunan, Guangxi, and Guizhou. Two cities and four counties were sampled in each province. Four neighborhoods (including suburban villages) in each city, and one county-town neighborhood and three villages in each county, were then randomly selected. Approximately 20 households were sampled per neighborhood or village.

The first survey was conducted in 1989. Follow-up surveys were conducted in 1991, 1993, 1997 and 2000. Since the over-time changes of child education attainment and most of the independent variables are very small, we focus on cross-sectional analysis drawing on the 1989 data. We do use the 1991 and 1993 data to check whether there is any change in terms of gender bias across birth cohorts and ages.[1]

This data set contains detailed information on characteristics of each individual living in the household, such as years of education, family position, as well as detailed

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[1] The 1997 and 2000 data are not perfectly compatible to early data.

household characteristics such as family income. It also contains a community survey which collects detailed community level information. We will draw on the birth control policy variable from the community survey for Chapter 4.

In this thesis, we refer to "children" as those that are between 6 and 15 years old, who are expected to be in school. In China, the educational system is a "6+3(+3+4)" scheme, which consists of six years of primary school, three year of junior secondary school, followed by three years of senior secondary or technical college, and four years of tertiary education.[2] Primary education generally starts at the age of 6 and junior secondary school ends at the age of 15. In 1986, the People's Congress of China passed the education law, which enforces 9-year compulsory education for children who reach the age of 6. However, the law has not been very effective in rural areas, where it is normal to see children have delayed enrolment and drop out of school at early ages. Following the literature (Rosenzweig and Wolpin, 1980; Subramaniam, 1996; Ota and Moffatt, 2002), we restrict the sample of children to be under 15 years old. We do so because these children are supposed to be in school and older children are more likely to move out of the household, which is not covered by the survey, and thus causes a sample selection problem. In total, we have 2417 observations for the analysis of gender bias and 1920 observations for examining the

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[2] Generally, most technical colleges offer three-year training, but some also offers 4-6 years of training which depends on the characteristics of the courses. This is also true for the tertiary education that some programs require 5 years to complete, such as Engineering and Architecture.

quantity-quality tradeoff. The descriptive statistics for the two samples are presented in Tables 1 and 2.

Our dependent variable, total years of education, refers to the number of years of formal education completed by the child. On average, the sampled children are 10.8 years old, who were expected to receive 4.8 years of education by 1989 (Table 1). However, they only received 4.2 years of education by 1989, which implies that on average, they have 0.6 years of delay in enrolment. The average years of education of fathers is 6.6 while that of mothers is 4.3. This shows that females from the last generation generally received 2.3 fewer years education than males. Among these 2,417 children, 21 percent of them are from urban area, 36 percent are the only child of their families, and 9 percent are from families with female household heads.

3 Gender and Birth Order of children

3.1 Hypothesis Development

3.1.1 Gender Discrimination

How do parents allocate household resources among children greatly depends on the value of each child to parents. The value varies with a child's gender and age, the parents' education levels, as well as other child and household characteristics (Ahn, 1995). When gender is a consideration, parents may very likely to allocate more resources to children of one gender, normally boys, than the other gender. We call this gender bias.

There is a huge empirical literature that attempts to measure intra-household gender bias among children in developing countries, but no consensus has been reached so far. Some find that gender bias exists in dimensions such as nutrition, health, and education (see e.g., Drèze and Sen, 1989; Harriss, 1990; Dasgupta, 1993; Strauss and Thomas, 1995; Ilahi, 2001; Kingdon, 2004). Others do not detect gender bias in their samples. For example, Subramaniam (1996) and Deaton (1998) use household level consumption data to analyze the intra-household allocation of resources. They find that the differential effects of gender composition in the allocation of resources derive entirely from the unobserved wealth effect, and the gender bias disappears after controlling for the fixed effect.

In this thesis, we focus on examining the gender bias in child education in China. In China, boys are very likely to be preferred by parents and thus receive more education for the following reasons. First, as an investment, a boy is for his parents' security in old age. This is especially true in rural areas, where there is no social security for old farmers, and where it is the son's responsibility to support and care for aging parents. From the parents' point of view, investing on boys will generate future income for the family for the whole life of parents, but investing on girls will only generate income until they are married to other families.[3] Second, in China, as well as in many developing countries, women's income is lower than men's, this will make investing in boys generate a higher return than in girls even if both share equally the responsibility of caring aging parents. Finally, a boy can carry on the family name, which is very important in the Chinese culture. Chinese families who do not have a son are discriminated against by friends and relatives because failure to carry on the family name is a serious sign of disrespect to ancestors.[4] Because of these reasons, boys have higher investment value than girls.

Our first hypothesis on gender bias is as follows:

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[3]In China, especially in rural areas, after marriage, boys will live with their own parents, while girls will live with her husband's parents.

[4]There is a saying in Chinese that describes this vividly: There are three disrespects we could have for our ancestors, not carrying on the family name is the biggest one ("bu xiao you san, wu hou wei da"). The discrimination can also be shown by two other Chinese sayings: "no sons, no grandsons" (duan zi jue sun), and "extinction of descendants" (jue hou), which are extremely negative sayings about a family.

Hypothesis 1: Among children with same age and similar family background, boys generally receive more education than girls.

To test Hypothesis 1, we estimate the following equation,

$$E_i = \alpha_i + \beta_1 G_i + \beta_2 A_i + \beta_3 A_i^2 + \beta_4 A_i G_i + X_i \beta_5 + \varepsilon_i \qquad (1)$$

where E_i is the total years of education of child i; G_i indicates the gender of the child which equals one for girls and zero for boys; A_i is the age of the child; and X_i is a vector of household variables including log household income per capita, parents' age and education, the female household head dummy, a urban dummy, and provincial dummies. The gender difference of education between boys and girls is $\beta_1 + \beta_4$*age.

In addition to examining gender bias across age groups, we are also interested in whether family characteristics, which may affect resource allocation on child education, have significant effects on gender bias. These characteristics include a dummy indicating whether a family has two or more children, household income per capita, the gender of household head, and parents' education levels.

Having more children may increase the gender bias. More children in a family mean less resource for each child on average. However, if boys are preferred, their resources may be reduced less than girls with an additional child in the family.

Therefore, more children may lead to stronger gender bias against girls.

Similarly, household income may also affect the size of gender bias. Lower income families face larger budget constraint in children education. Therefore, if boys are given higher priority in education than girls, girls from lower income families may receive less education than girls from higher income families, which means the size of gender bias may decrease with household income.

Both the status of the mother and parental education levels are important in affecting resource allocation within a family. Prior research has found mothers are more likely to allocate more resources to girls, and thus the status and education of the mother is very important in reducing the gender bias against girls. For example, by using the data from the United States, Ghana, and Brazil, Thomas (1994) finds that mother's education has a larger effect on the daughter's height while father's education affects the son's height more. This implies that mothers allocate more resources to their daughters while fathers channel more resources towards their sons.

To estimate the effect of these family characteristics on the size of gender bias, we have Hypothesis 2 as follows.

Hypothesis 2: The education difference between boys and girls is smaller when the mother is the household head or has more education, and when household income is *higher; it is larger when the father has more education and there are more children in the households.*

To test Hypothesis 2, the following equation is set up by adding some interactive terms to equation (1):

$$E_i = \alpha_i + \beta_1 G_i + \beta_2 A_i + \beta_3 A_i^2 + \beta_4 A_i G_i + \beta_5 H_i + \beta_6 H_i G_i + X_i \beta_7 + \varepsilon_i \qquad (2)$$

where H_i represents the household characteristics including the two-or-more children dummy, log household income per capita, the household head gender, and parents' education levels, and $H_i\ G_i$ represents the interactive terms of gender with these variables. If hypothesis 2 is valid, β_6 should be significantly different from zero.

3.1.2 Birth Order

Besides the gender effect, the birth order of children may also affect resource allocation, and the direction of the effect is ambiguous. Psychologists and sociologists suggest that first-born children have better innate abilities than their later-born siblings (Zajonc 1976). Parents may then allocate more resources to first-born children as their expected return to education is higher. On the other hand, if borrowing is limited and income stream is upward sloping, parents will be financially better to raise later-born children (Kessler 1991).

Behrman and Taubman (1986) use data from U.S. to test for birth-order effects on schooling and earnings. After controlling for family background, they find significant birth-order effects favoring firstborns on schooling opportunities. Rubalcava and Contreras (2000) use Chilean children's nutritional status to test whether birth-order in family and whether being a son or being a daughter reflect how parents allocate resources. They find the same result as Thomas (1994) that, mothers give more resources to their daughters and fathers to their sons, but the effect is only significant for non-oldest daughters and non-oldest sons.

Apart from the pure birth-order effect, we are also concerned about whether the birth order affects gender bias. For example, it is expected that first child will receive better education on the average, but is there any significant different between male first child and female first child? In addition to the significant gender bias against girls on school enrollment and drop-out rates, Ota and Moffatt (2002) find that first-borns are significantly less likely to be educated than the other children, and this disadvantage appears to apply equally to both genders. Besides, children are more likely to go to school if their elder siblings are girls, and less likely if their younger siblings are boys, which means older sisters are the most disadvantaged.

To test for the birth-order effect, hypothesis 3 is set up:

Hypothesis 3: Children's positions among siblings have significant effect on their education. The birth order may affect the magnitude of the gender effect.

The following equation is set up for testing hypothesis 3,

$$E_i = \alpha_i + \beta_1 G_i + \beta_2 A_i + \beta_3 A_i^2 + \beta_4 A_i G_i + \beta_5 O_i + \beta_6 O_i G_i + X_i \beta_7 + \varepsilon_i, \quad (3)$$

where O_i is a vector of birth order dummies indicating whether the child is the first or last child of a family. The coefficients β_5 and β_6 measure the birth order effect as well as the gender-birth order interaction effect.

3.2 Results on Gender Bias

In this section, we test Hypotheses 1 and 2 drawing on the CNHS data. Although we have a panel dataset, since the over-time changes of child education attainment and most of the independent variables are very small, we focus on cross-sectional analysis drawing on the 1989 data. However, we also use the 1991 and 1993 data to check whether there is any change in terms of gender bias across birth cohorts and ages.

3.2.1 Overall Results

In the first column of Table 3, we report the Ordinary Least Squares (OLS) estimate of the determinants of child educational attainment, where the dependent variable is the years of schooling of a child. The independent variables are child gender, age, age squared, log per capita household income, father's age, mother's age, father's education, mother's education, the female household head dummy, the urban dummy and provincial dummies. Note that the coefficient of child gender is not significantly different from zero, which means that on average there is no gender bias on child education.

However, after adding the interaction term of age and gender in regression (2), we find a gender bias against older girls. Note that the coefficients of both the gender dummy and the interaction term are significant, implying a varying gender effect across age. In Table 4, we report the estimated effect of being a girl on education for each age group, which is $\beta_1+\beta_4$*age. According to the first two columns of Table 4, a six-year-old girl generally receives 0.27 years more education than a six-year-old boy. However, the advantage of being a girl on education decreases with age and becomes negative at the age of 10. For the oldest children in our sample, i.e., those are fifteen-year-old, girls receive 0.36 fewer years of education than boys. To conclude, gender bias against girls is more serious for older children. Therefore, Hypothesis 1 is supported for older children.

There could be two ways to interpret the age-varying gender effect. First, although there is no bias against younger girls in terms of education, there is a bias against older girls. Second, only girls of the older birth cohorts were discriminated against, there is no bias against girls for the younger cohorts. One way to differentiate the two interpretations is to look at the education of the same cohorts of children when they become older. To do this, we apply the same regression models to the 1991 and 1993 samples, which include children in the 1989 sample, who were two and four years older in 1991 and 1993 correspondingly (but no older than 15), as well as some new children, who have reached the age of 6 by 1991 and 1993 respectively.

Regressions using the 1991 and 1993 samples suggest that the gender bias does not exist for samples of children that are of the same age but from younger birth cohorts. In columns 3 and 6, the coefficients of gender and age*gender are not significant. Nonetheless, we calculate the gender bias for each age group according to the estimated coefficients on age and age*gender even they are insignificant. The last four columns of Table 4, which report the gender bias by age for the 1991 and 1993 samples, show that there is almost no gender bias. For the 1991 sample, the calculated gender biases are all positive, suggesting that girls receive more education than boys for each age group. For the 1993 sample, only girls at 13 or older ages receive less education than boys, but the magnitudes of these biases are very small compared to the gender biases for the same age groups in 1989. These findings suggest that gender bias may not be age-specific, but is cohort-specific, i.e., gender bias only exists for older birth cohorts. In other words, gender bias is becoming less serious over time.

Most of the control variables have expected signs in these regressions. The years of education increases with age, but at a decreasing rate, with the age having a positive coefficient while age squared having a negative coefficient. Children in richer household and living in urban areas, where more government resources are allocated, have more education than their counterparts. Although the ages of both parents are not important for child education, the education levels of both parents are important determinants of child education. A little unlike the evidence from other developing countries (Alderman et al., 2001; Glewwe and Jacoby, 1995), we find that, although the education of both parents has positive and significant coefficients, that on the father's education is larger. This probably reflects the patrilineality that exists within Chinese society, where fathers make major household decisions, including child education.[5] Another interesting finding is that children in female-headed households

[5] This is consistent with the findings of Brown and Park (2002) and Li et al. (2005).

on average have 0.334 years of more education than children in male-headed households.

3.2.2 Gender Bias in Rural versus Urban Areas

In the above analysis, we have assumed that the gender bias is the same in rural and urban areas. However, the education system and the return to education for boys versus girls may be very different in rural and urban areas. Since rural families have less social security and investment opportunities, they reply more on having boys for old-age security. As a result, rural families may have a relative larger return in investing in boys relative to girls, and thus a larger gender bias.

To analyze whether there is a significant rural and urban difference, we do the same estimations to the rural and urban samples separately. Results presented in the first two columns of Table 5 show that most of the results of gender bias found in the 1989 sample (in Table 3) are coming from rural families. Note that the coefficients of gender and age*gender interaction are both significant for the rural sample, but they are not significant for the urban sample. This shows that there is no detectable gender bias in the urban area, and any gender effect found in the 1989 sample is totally due to the gender effect in rural China. The rural-urban difference suggests that we need to separate the rural and urban samples in the following analysis. When we do the same regressions to the 1991 and 1993 sample, there is no detectable gender bias in both the rural and urban samples, suggesting that gender bias indeed has become unimportant for children of the younger birth cohorts.

3.2.3 Household Characteristics and Gender Bias

We next test Hypothesis 2, i.e., whether the gender bias changes with other household characteristics such as the number of children, the gender of the household head and the education levels of both parents. We first carry out the analysis using the rural sample, and results are reported in Table 6. In column 1, we include both the two-or-more children dummy as well as its interaction with gender. The interaction term is negative but not significant, which implies that having more than one child does not have a significant effect on the magnitude of gender bias in rural China.

The other four household variables indeed affect gender bias in education in rural China. In column 2 of Table 6, we include the interactive term of the female household head dummy and gender. Interestingly, the female household head dummy becomes insignificant while the interactive term is significantly positive. The insignificant female household head dummy means that for boys, there is no significant difference in education between female-headed and male-headed families. The significant interaction term means that girls from female-headed families received more education than girls from male-headed families. Thus, the significantly positive effect of female household head on child education found in previous regressions (in Tables 3 and 5) is totally due to more education received by girls from female-headed families. Our results are similar to those of Thomas (1994) in that female household heads direct more resources to girls, but also different in the sense that, male household heads do not necessarily direct more resources to boys.[6] One reason to explain for the female household head effect on gender bias is that, women are less concerned than men with the perpetuation of ancestral lineage, which is the main cause of gender bias against girls in China.

Regressions 3 to 5 in Table 6 show that the education of parents, and in particular that of the father, affects the magnitude of the gender bias in rural China. In these columns, we add the interaction terms of mother's and father's education with gender. With these interaction terms added, both the mother's and father's education levels continue to have a positive effect, suggesting that parental education has a positive effect on the education of boys. However, the father's education has a larger coefficient and is also more significant. Moreover, the coefficients of the interaction terms are all positive, and are consistently significant for that of the father, suggesting that parental education and in particular the father's education is more important for

[6] If male household heads direct more resources to boys, the coefficient of female household head dummy in regression (2) of Table 6 should be significantly negative.

the education of girls, though the effect is very small.[7] Our finding is again different from that of Thomas (1994), who finds that mother's education has a larger effect on her daughter's achievement while father's education has a larger effect on his son's achievement. Our finding seems to suggest the father's education is more important in increasing the education of daughters. Although men are generally more concerned with the perpetuation of ancestral lineage, education can promote the importance of gender equality, which accounts for the result of having a better educated father increases girls' education.

In the last column of Table 6, the interactive term of log household income per capita and gender is included. Similar to the effect of female household head, log household income per capita becomes insignificant while the interactive term is significantly positive, which means the level of household income does not have significant effect on boys' education, but have significantly positive effect on girls' education. In other words, higher household income decreases the size of gender bias. This result provides a reason for the disappearing of gender bias over time, which is caused by the increasing average household income through the process of economic development.

We carry out the same analysis using the urban sample, and the results are

[7] The coefficients of both interactions are about 0.02, which means when father's or mother's education increased by one year, their daughter's education will be increased by 0.02 year.

presented in Table 7. We find that neither the gender dummy itself nor its interactions with child age and other household characteristics are significant. Our results consistently show that there is no gender bias in education in urban China.

To summarize, the above results show that gender bias against girls only exist in rural China and only for children of the older birth cohorts. For the younger cohorts of rural children, girls even have more education than boys. We also find that for the rural sample, female-headed households tend to favor the education of girls, and both the education level of fathers and higher household income per capita help to reduce the gender bias. However, we do not find that having the second or more children changes the gender bias.

3.3 Birth Order Effect

In this section, we test Hypothesis 3 regarding the effect of birth order on gender bias. Our measures of birth order include the first and last child dummies, which refer to the eldest and youngest child respectively in families with two or more children. Therefore, children with the two dummies equals to 0, as well as the two-or-more children dummy equals to 1, represents those middle-order children from households with three or more children.

We first present results using the rural sample in Table 8. The result of regression

(1) in Table 8 shows that the two birth order dummies, the first child and last child dummies, are positive and significant in rural China in 1989, while the two-or-more children dummy is significantly negative. Being the first and last child in a family with two or more children increases education by 0.356 and 0.193 years respectively, being the only child increases education by 0.241 years.[8] These numbers indicate that the average education of children in two-child families (0.275) is even larger than the education of children from one-child families. In addition, being the first child in a large family is more advantageous in terms of education than being the only child.

Our further analysis shows that the gender bias does not change with the birth order dummies. In regression (2) and (3), we include interactions of gender with these birth order dummies. None of these newly added interaction terms are significantly different from zero, suggesting that birth order does not affect gender bias in rural China.

The results regarding the urban sample are generally the same as those of the rural sample. As shown by regressions in Table 9, being the only child, first child and last child all increase education of urban children, though only the effects on the first and last child are statistically significant. Interestingly, being the only child in urban China does not increase education significantly. Again, none of the interaction terms

[8] The effect of being the only child is the same but in opposite direction with the coefficient of two or more children dummy.

of gender with birth order are significant.

4 Number of children

4.1 Hypothesis and Empirical Strategy

After investigating the effect of children's gender and birth order on the education level they achieved, we are now going to analyze the effect of total number of children on child education. The quantity-quality model was first introduced in Becker (1960) and expanded in Becker and Lewis (1973), which suggests that the interaction between quantity and quality in the budget constraint leads to rising marginal cost of quality with respect to family size, therefore caused the tradeoff between quantity and quality of children.

To clearly identify this effect, the problem of endogeneity between family size and resource allocation must be solved. Rosenzweig and Wolpin (1980) are the first to use twins as a natural experiment to identify the exogenous increases in family sizes in India. They confirm that the exogenous increases in fertility reduce child quality, which implies that exogenous improvements in birth control technology would increase schooling level of Indian children. Black, Devereux and Salvanes (2004) use a richer data set in Norway to do a similar research. The birth of twins is also used as instrumental variables to isolate causation, and they find that the negative correlation between family size and children's education attainment is only significant when birth order is not controlled. In addition, they find that children born later in the family obtain less education.

Instead of using education attainment, Lee (2004) uses parents' monetary investment in children's education to measure children quality. He uses the first child's sex as instrument for fertility to test for quantity quality tradeoff in South Korea. The results show that tradeoff still exists after removing endogeneity, although not as strong as that estimated by ordinary least square regressions.

Qian (2004) exploits the exogenous changes in family size caused by relaxation of family planning policy in China to estimate the effect of family size on school enrollment. She finds that school enrollment for girls from one-child families increased by 18 to 20 percent when parents had an additional child, which implies that quality is not monotonically decreasing in quantity.

To analyze this tradeoff issue in China, hypothesis 3 is set up.

Hypothesis 4: There is Quantity-Quality tradeoff of children, which means children from larger family generally receive less education than children from small families.

To clearly identify the relationship between the number of children and the education attained by each child, the problem of endogeneity must be solved. Endogeneity exists if both the independent variable and the dependent variable are correlated with some unobserved or uncontrolled variables. This makes the causal relationship between independent and dependent variables ambiguous. As both the number of children and education resources allocated to each child are household choice variables which are correlated with unobserved household preferences, the effect of endogeneity cannot be ignored.

To remove endogeneity, the method of instrumental variables (IVs) is used in this study. A good instrumental variable should be highly correlated with the independent variables, but not correlated with the dependent variable except through the effect of the independent variable. All other household choice variables are not good IVs to fertility as they still commit certain level of endogeneity. Twins birth (Rosenzweig and Wolpin 1980; Black et. al. 2004) and the sex composition of children (Lee 2004; Conley 2004) are two commonly used instruments of fertility in previous literature. However, studies using twins data are usually constrained by limited data, and sex composition may still be a choice variable given that abortion is not illegal.

In this thesis, community level variable concerning the birth control policy is used.[9] Under the assumption that the birth control policy only affects the number of children but not directly affects child education, it can serve as an IV. Qian (2004) also uses the one-child policy as a natural experiment to detect the effect of

[9] In our data, a community refers to a town, a village, or a neighborhood.

exogenous increase in children quantity on children quality. She uses the triple interaction term of an individual's sex, date of birth and origin of birth as instrument to capture the effect of relaxation of one-child policy in a particular group of children, which is positively correlated to the fertility of the affected households. In analyzing the quantity quality tradeoff, this study is different from her study in two ways. First, Qian (2004) uses the individual data to estimate the effect of increasing fertility on individual child, while the household data is used in this study to test whether there is quantity quality tradeoff on the average education of all children in the same household. Second, while Qian uses the different effects of policy on different groups to capture the exogenous change in family size, community variable of one-child policy is directly used as instrumental variable in this thesis.

One-child policy was formally announced in 1979 in China, which only allowed each household to have one child in order to control the rapid population growth rate. Under this country-level policy, each province or even each community can still have some variations on their own policies, depending on their population structures. These community-level policy variables are good instruments to fertility as they are highly correlated with fertility but not directly correlated with child education. For example, if the birth-control policy is more flexible in a community, families in this community will have more children on average, but a more flexible birth-control policy will not have any direct effect on the resources distributed to child education by the parents. Among these policy variables, we found that the size of fine for an extra-child (F_i) has the largest correlation with the two or more children dummy in the first-stage of the two-staged least squares regressions, and is then used as an instrument for fertility in this thesis. With heavier punishment for extra child, the average fertility of families in the community will be lower.

The following equation is first estimated by the Ordinary Least Squares regressions (OLS) followed by the two-staged least squares (2SLS) regressions to test for the validity of hypothesis 4:

$$e_i = \alpha_i + \beta_1 N_i + X_i\beta_2 + \varepsilon_i \quad (4)$$

where e_i is the years of education of children i; N_i is a dummy which equals to 1 if the family has two or more children; and X_i is a vector of other exogenous variables including child gender, age and its squared term, parents' ages and education levels, and a dummy indicating whether the family is a urban household. N_i measures the quantity while e_i measures the quality of children in the household. If quantity quality tradeoff exists, β_1 should be smaller than zero.

4.2 Results

Equation (4) is first estimated by OLS and the results are presented in the first column of Table 10. The results show that β_1 is significantly different from zero in 1989, which represent there is tradeoff between number of children and the mean education of all children in the household. However, this result may be biased due to the problem of endogeneity discussed before. The OLS results suggest that total number of children in the household is negatively correlated to child education, but it can also be the case that families plan to invest less in child education, also prefer to have more children due to some unknown factors. If that is the case, the negative impact of children quantity on children quality estimated by OLS will be over-estimated. To remove endogeneity and identify the true effect of children quantity on children quality, 2SLS is used instead.

The first stage of the 2SLS regressions estimates the correlation between fertility and the instruments:

$$N_i = \alpha_i + \beta_1 F_i + X_i \beta_2 + \varepsilon_i \tag{5}$$

where F_i is the instrument to total number of children and X_i is the same vector as equation (4). The result is presented in Table 11. F_i is found to be significant and negatively correlated with N_i, which is the same as our expectation. Recall that a good instrument should be highly correlated with the independent variable but not directly correlated with the dependent variable. The first-stage of 2SLS regressions shows that F_i is highly correlated with fertility, and it is obvious that individual household's choice on child education will not affect the birth-control policy of the community. Therefore, it is suitable to use F_i as IV for fertility.

The second column of Table 10 shows the second stage results of the two-stage least squares regressions of equation (4). After removing endogeneity by the instrumental variables approach, we can still find significant tradeoff between quantity and education level of children, and the negative effect is even 1.28 years larger than that estimated by OLS, which means children from one-child families got one year more education than children from families with two or more children on the average. This implies that the OLS estimates are actually positively biased due to the problem of endogeneity, that is, families that planned to have more children also planned to educate their children more due to unobserved or omitted variables.

In China, most parents treat children as investment for old-age security. These parents may choose to have more children in order to diversify risk, so that they will still have old-age support if one of the children is not able to earn money due to health or other constraints. At the same time, these parents will invest a large proportion of resources in child education as they believe that education level is positively correlated with the children's future income. With this unobserved preference of parents, the correlation between number and education of children estimated by simple OLS regressions will be positively biased, resulting in a smaller or even no quantity-quality tradeoff. After removing endogeneity by IV method, the results of 2SLS regressions reflect the true size of quantity and quality tradeoff: the education levels of children in one-child families are 1.37 year more than those at the same ages but from families with two or more children in 1989.

5 Conclusions

In this thesis, we have studied the child education attainment in China. We first examine the existence of gender bias in China. We find that gender bias against girls only exist in rural China and only for children of the older birth cohorts. For the younger cohorts of rural children, girls even have more education than boys.

In addition to examining gender bias across age groups, we are also interested in whether family characteristics and child birth order, which may affect resource allocation on child education, have significant effects on gender bias. We find that for the rural sample, female-headed households tend to favor the education of girls, and both the education level of fathers and higher household income per capita help to reduce the gender bias. However, we do not find that having one more child or the mother's education affects the magnitude of gender bias. Our further analysis shows that although both the first and last births in a family receive more education than their siblings, girls of the first and last births are neither favored nor discriminated in terms of education.

Finally, we examine whether there is a quantity-quality tradeoff of children. By directly employing the one-child policy variable as an instrumental variable, we find that there is indeed a quantity-quality tradeoff of children. Further analysis that draws on even better instrumental variables is needed in the future.

Our analysis has some policy implications. Our empirical findings seem to suggest that the gender bias has become less serious and even disappearing over time. Thus, no specific policy is needed to raise the education level of girls. Since there is a quantity-quality tradeoff, the one-child policy of China may indeed have helped to raise the average quality of children and thus their long-term earnings ability and the growth of the Chinese economy.

Table 1: Descriptive Statistics of Children's Data Used in Section 3 (1989)

Variable	Number of observations	Mean	Standard Deviation	Minimum Value	Maximum Value
Years of education	2417	4.1734	2.6475	0	11
Gender of child (boy=0, girl=1)	2417	0.4820	0.4998	0	1
Age of child	2417	10.7605	2.8799	6	15
Per capita income	2417	971.5823	790.5533	4.9642	8231.451
Father's age	2417	40.2263	6.5211	24	68
Mother's age	2417	38.1382	5.7943	22	62
Father's education level	2417	6.5742	3.4774	0	18
Mother's education level	2417	4.2913	3.8726	0	17
Female household head dummy (male=0, female=1)	2417	0.0910	0.2877	0	1
Urban dummy	2417	0.2110	0.4081	0	1
Two or more children dummy (two or more=1, one=0)	2417	0.6396	0.4802	0	1
First child dummy	2417	0.2065	0.40481	0	1
Last child dummy	2417	0.2350	0.4241	0	1

Table 2: Descriptive Statistics of Children's Data Used in Section 4 (1989)

Variable	Number of observations	Mean	Standard Deviation	Maximum Value	Minimum Value
Years of education	1920	4.193229	2.674773	0	11
Gender of child (boy=0, girl=1)	1920	0.484375	0.499886	0	1
Age of child	1920	10.43698	2.797152	6	15
Father's age	1920	39.76094	6.45614	26	68
Mother's age	1920	37.71719	5.754733	22	62
Father's education level	1920	6.600521	3.358282	0	18
Mother's education level	1920	4.216146	3.830658	0	17
Urban dummy	1920	0.173958	0.379172	0	1
Two or more children dummy (two or more=1, one=0)	1920	0.671875	0.469653	0	1
Amount of fine for extra child	1920	1506.293	1531.127	0	9999

Table 3: Ordinary Least Squares Regressions Examining the Effect of Gender on Child's Education in China

	Dependent variable: Total Years of Education					
	1989		1991		1993	
	(1)	(2)	(3)	(4)	(5)	(6)
Gender of child	-0.057	0.692***	0.056	-0.014	0.062	0.369*
(boy=0, girl=1)	(1.04)	(3.27)	(1.04)	(0.07)	(1.19)	(1.90)
Age of child	1.165***	1.196***	0.939***	0.935***	0.830***	0.839***
	(14.26)	(14.60)	(11.76)	(11.64)	(11.00)	(11.10)
Age squared	-0.018***	-0.018***	-0.006	-0.006	0.001	0.001
	(4.76)	(4.75)	(1.51)	(1.51)	(0.29)	(0.36)
Age*gender		-0.070***		0.007		-0.029
		(3.67)		(0.35)		(1.64)
Log income per capita	0.072**	0.073**	0.140***	0.140***	0.128***	0.127***
	(2.17)	(2.20)	(3.94)	(3.94)	(4.78)	(4.76)
Father's age	0.012	0.011	-0.001	-0.001	0.012	0.012
	(1.35)	(1.30)	(0.13)	(0.14)	(1.41)	(1.42)
Mother's age	0.005	0.005	0.009	0.009	-0.004	-0.005
	(0.50)	(0.56)	(0.84)	(0.84)	(0.47)	(0.48)
Father's education level	0.078***	0.078***	0.047***	0.047***	0.032***	0.031***
	(8.23)	(8.21)	(4.86)	(4.87)	(3.41)	(3.40)
Mother's education level	0.049***	0.049***	0.027***	0.027***	0.048***	0.048***
	(5.32)	(5.33)	(3.04)	(3.04)	(5.68)	(5.73)
Female household head	0.330***	0.334***	-0.054	-0.054	0.219**	0.216**
(male=0, female=1)	(3.37)	(3.41)	(0.50)	(0.50)	(2.14)	(2.10)
Urban dummy	0.330***	0.330***	0.368***	0.368***	0.358***	0.362***
	(4.48)	(4.50)	(5.08)	(5.08)	(5.18)	(5.24)
Observations	2417	2417	2229	2229	2123	2123
R-squared	0.74	0.75	0.77	0.77	0.82	0.82

Numbers in parentheses are absolute values of t statistics. Significance levels of 0.1, 0.05 and 0.01 are noted by *, **, and ***. All regressions include provincial dummies.

Table 4: Gender Effects of Different-year-old Girls in 1989, 1991 and 1993

Year of birth	1989		1991		1993	
	Age	Gender bias	Age	Gender bias	Age	Gender bias
1987					6	0.195
1986					7	0.166
1985			6	0.028	8	0.137
1984			7	0.035	9	0.108
1983	6	0.272	8	0.042	10	0.079
1982	7	0.202	9	0.049	11	0.05
1981	8	0.132	10	0.056	12	0.021
1980	9	0.062	11	0.063	13	-0.008
1979	10	-0.008	12	0.07	14	-0.037
1978	11	-0.078	13	0.077	15	-0.066
1977	12	-0.148	14	0.084		
1976	13	-0.218	15	0.091		
1975	14	-0.288				
1974	15	-0.358				

Table 5: Ordinary Least Squares Regressions Examining the Effect of Gender on Child's Education in China

	Dependent variable: Total Years of Education					
	1989		1991		1993	
	Rural	Urban	Rural	Urban	Rural	Urban
Gender of child	0.873***	0.076	-0.156	0.549	0.341	0.491
(boy=0, girl=1)	(3.56)	(0.20)	(0.65)	(1.39)	(1.55)	(1.19)
Age of child	1.218***	1.096***	0.904***	1.061***	0.889***	0.679***
	(12.79)	(7.72)	(9.80)	(7.02)	(10.35)	(4.29)
Age squared	-0.020***	-0.010	-0.005	-0.006	-0.001	0.009
	(4.57)	(1.46)	(1.26)	(0.88)	(0.35)	(1.26)
Age*gender	-0.088***	-0.006	0.016	-0.032	-0.032	-0.024
	(4.01)	(0.18)	(0.73)	(0.90)	(1.58)	(0.64)
Log income per capita	0.084**	-0.078	0.169***	-0.120	0.144***	0.024
	(2.32)	(0.91)	(4.35)	(1.36)	(4.99)	(0.32)
Father's age	0.011	-0.009	-0.001	-0.011	-0.001	0.055***
	(1.06)	(0.65)	(0.07)	(0.72)	(0.11)	(3.08)
Mother's age	0.000	0.051***	0.006	0.037**	0.003	-0.009
	(0.02)	(3.03)	(0.51)	(2.03)	(0.27)	(0.44)
Father's education level	0.091***	0.032**	0.058***	0.010	0.033***	0.026
	(7.94)	(2.15)	(5.02)	(0.60)	(3.03)	(1.49)
Mother's education level	0.041***	0.081***	0.024**	0.053***	0.048***	0.050***
	(3.73)	(5.53)	(2.36)	(3.25)	(4.92)	(2.89)
Female household head	0.312**	0.195	-0.006	-0.111	0.272**	0.049
(male=0, female=1)	(2.48)	(1.44)	(0.05)	(0.75)	(2.18)	(0.28)
Observations	1907	510	1781	448	1683	440
R-squared	0.71	0.87	0.75	0.87	0.80	0.86

Numbers in parentheses are absolute values of t statistics. Significance levels of 0.1, 0.05 and 0.01 are noted by *, **, and ***. All regressions include provincial dummies.

Table 6: Ordinary Least Squares Regressions Examining the Effect of Gender on Child's Education Rural Area of China (1989)

	Dependent variable: Total Years of Education					
	(1)	(2)	(3)	(4)	(5)	(6)
Gender of child	0.978***	0.852***	0.351	0.558**	0.279	-0.028
(boy=0, girl=1)	(3.78)	(3.48)	(1.19)	(2.05)	(0.93)	(0.06)
Age of child	1.238***	1.217***	1.215***	1.205***	1.207***	1.214***
	(12.75)	(12.78)	(12.78)	(12.65)	(12.69)	(12.76)
Age squared	-0.021***	-0.020***	-0.020***	-0.020***	-0.020***	-0.020***
	(4.68)	(4.54)	(4.59)	(4.49)	(4.54)	(4.53)
Age*gender	-0.085***	-0.089***	-0.075***	-0.075***	-0.070***	-0.088***
	(3.87)	(4.05)	(3.35)	(3.36)	(3.11)	(4.00)
Log income per capita	0.081**	0.083**	0.081**	0.083**	0.081**	0.019
	(2.22)	(2.29)	(2.25)	(2.29)	(2.25)	(0.39)
Father's age	0.091***	0.092***	0.061***	0.091***	0.068***	0.090***
	(8.00)	(8.02)	(4.15)	(7.97)	(4.39)	(7.92)
Mother's age	0.040***	0.041***	0.041***	0.018	0.027*	0.041***
	(3.57)	(3.74)	(3.74)	(1.26)	(1.85)	(3.70)
Father's education level	0.010	0.010	0.010	0.011	0.010	0.011
	(1.01)	(0.99)	(0.96)	(1.04)	(0.97)	(1.02)
Mother's education level	-0.000	0.001	0.000	0.000	0.000	0.001
	(0.00)	(0.12)	(0.04)	(0.04)	(0.04)	(0.05)
Female household head	0.305**	0.095	0.330***	0.313**	0.326***	0.305**
(male=0, female=1)	(2.42)	(0.55)	(2.63)	(2.49)	(2.60)	(2.43)
Two or more children dummy	0.025					
(two or more=1, one=0)	(0.26)					
Two or more children dummy	-0.192					
*gender	(1.41)					
Female household head dummy		0.446*				
*gender		(1.81)				
Father's education level			0.062***		0.048**	
*gender			(3.12)		(2.18)	
Mother's education level				0.048***	0.029	
*gender				(2.65)	(1.42)	
Log income per capita						0.139**
*gender						(2.03)
Observations	1907	1907	1907	1907	1907	1907
R-squared	0.71	0.71	0.72	0.72	0.72	0.71

Numbers in parentheses are absolute values of t statistics. Significance levels of 0.1, 0.05 and 0.01 are noted by *, **, and ***. All regressions include provincial dummies.

Table 7: Ordinary Least Squares Regressions Examining the Effect of Gender on Child's Education Urban Area of China (1989)

	Dependent variable: Total Years of Education					
	(1)	(2)	(3)	(4)	(5)	(6)
Gender of child (boy=0, girl=1)	0.127 (0.34)	0.082 (0.22)	-0.196 (0.44)	-0.192 (0.44)	-0.275 (0.60)	-1.456 (1.27)
Age of child	1.078*** (7.50)	1.109*** (7.81)	1.094*** (7.71)	1.092*** (7.69)	1.092*** (7.69)	1.102*** (7.77)
Age squared	-0.009 (1.34)	-0.010 (1.55)	-0.010 (1.46)	-0.010 (1.45)	-0.010 (1.45)	-0.010 (1.52)
Age*gender	-0.003 (0.09)	-0.012 (0.37)	-0.001 (0.04)	0.002 (0.07)	0.003 (0.08)	-0.009 (0.25)
Log income per capita	-0.067 (0.77)	-0.081 (0.93)	-0.086 (1.00)	-0.084 (0.97)	-0.087 (1.00)	-0.189 (1.62)
Father's age	0.032** (2.13)	0.033** (2.17)	0.020 (1.02)	0.033** (2.18)	0.025 (1.22)	0.031** (2.08)
Mother's age	0.082*** (5.58)	0.081*** (5.54)	0.082*** (5.58)	0.068*** (3.76)	0.072*** (3.66)	0.081*** (5.55)
Father's education level	-0.010 (0.71)	-0.009 (0.64)	-0.008 (0.62)	-0.009 (0.68)	-0.009 (0.65)	-0.009 (0.65)
Mother's education level	0.052*** (3.03)	0.052*** (3.06)	0.051*** (2.99)	0.051*** (3.02)	0.051*** (3.00)	0.052*** (3.07)
Female household head (male=0, female=1)	0.210 (1.53)	0.017 (0.09)	0.196 (1.44)	0.198 (1.46)	0.198 (1.45)	0.194 (1.43)
Two or more children dummy (two or more=1, one=0)	0.174 (1.23)					
Two or more children dummy *gender	-0.187 (0.95)					
Female household head dummy *gender		0.373 (1.44)				
Father's education level *gender			0.027 (1.07)		0.016 (0.52)	
Mother's education level *gender				0.027 (1.18)	0.020 (0.72)	
Log income per capita *gender						0.224 (1.41)
Observations	510	510	510	510	510	510
R-squared	0.87	0.87	0.87	0.87	0.87	0.87

Numbers in parentheses are absolute values of t statistics. Significance levels of 0.1, 0.05 and 0.01 are noted by *, **, and ***. All regressions include provincial dummies.

Table 8: Ordinary Least Squares Regressions Examining the Effect of Gender on Child's Education in Rural Area of China (1989)

	Dependent variable: Total Years of Education		
	(1)	(2)	(3)
Gender of child (boy=0, girl=1)	0.828*** (3.37)	0.815*** (3.31)	0.837*** (3.14)
Age of child	1.187*** (12.14)	1.189*** (12.17)	1.187*** (12.14)
Age squared	-0.019*** (4.22)	-0.019*** (4.27)	-0.019*** (4.22)
Age*gender	-0.082*** (3.75)	-0.078*** (3.47)	-0.083*** (3.63)
Log income per capita	0.077** (2.11)	0.075** (2.06)	0.077** (2.11)
Father's age	0.012 (1.13)	0.012 (1.12)	0.012 (1.13)
Mother's age	0.004 (0.33)	0.004 (0.35)	0.004 (0.33)
Father's education level	0.093*** (8.16)	0.093*** (8.15)	0.093*** (8.15)
Mother's education level	0.037*** (3.37)	0.037*** (3.35)	0.037*** (3.37)
Female household head (male=0, female=1)	0.313** (2.49)	0.312** (2.48)	0.313** (2.49)
Two or more children dummy (two or more=1, one=0)	-0.241*** (2.59)	-0.248*** (2.66)	-0.241*** (2.59)
First child dummy	0.356*** (3.59)	0.452*** (3.45)	0.356*** (3.59)
Last child dummy	0.193* (1.93)	0.199** (1.98)	0.199* (1.65)
First child dummy*gender		-0.175 (1.12)	
Last child dummy*gender			-0.014 (0.09)
Observations	1907	1907	1907
R-squared	0.72	0.72	0.72

Numbers in parentheses are absolute values of t statistics. Significance levels of 0.1, 0.05 and 0.01 are noted by *, **, and ***. All regressions include provincial dummies.

Table 9: Ordinary Least Squares Regressions Examining the Effect of Gender on Child's Education in Urban Area of China (1989)

	Dependent variable: Total Years of Education		
	(1)	(2)	(3)
Gender of child (boy=0, girl=1)	0.096 (0.26)	0.115 (0.31)	0.146 (0.38)
Age of child	1.027*** (7.10)	1.026*** (7.08)	1.028*** (7.10)
Age squared	-0.007 (1.01)	-0.007 (0.98)	-0.007 (0.99)
Age*gender	-0.007 (0.22)	-0.011 (0.32)	-0.010 (0.28)
Log income per capita	-0.083 (0.95)	-0.085 (0.98)	-0.084 (0.97)
Father's age	-0.009 (0.65)	-0.009 (0.68)	-0.009 (0.69)
Mother's age	0.056*** (3.26)	0.057*** (3.28)	0.056*** (3.25)
Father's education level	0.032** (2.11)	0.032** (2.12)	0.032** (2.10)
Mother's education level	0.082*** (5.55)	0.081*** (5.54)	0.082*** (5.55)
Female household head (male=0, female=1)	0.205 (1.51)	0.203 (1.49)	0.207 (1.52)
Two or more children dummy (two or more=1, one=0)	-0.253 (1.42)	-0.249 (1.39)	-0.258 (1.44)
First child dummy	0.493** (2.48)	0.422* (1.66)	0.492** (2.47)
Last child dummy	0.370* (1.87)	0.366* (1.86)	0.429* (1.87)
First child dummy*gender		0.120 (0.45)	
Last child dummy*gender			-0.126 (0.51)
Observations	510	510	510
R-squared	0.87	0.87	0.87

Numbers in parentheses are absolute values of t statistics. Significance levels of 0.1, 0.05 and 0.01 are noted by *, **, and ***. All regressions include provincial dummies.

Table 10: Ordinary Least Squares and Two-Stage Least Squares Regressions Examining the Effect of Number of Children on Education of Children in China (1989)

	Dependent variable: Mean Years of Education	
	OLS	2SLS
Two or more children dummy (two or more=1, one=0)	-0.092** (1.97)	-1.370*** (3.29)
Gender of child (boy=0, girl=1)	0.103** (2.54)	0.115** (2.38)
Age of child	0.669*** (10.76)	1.011*** (7.62)
Age squared	0.010*** (3.44)	-0.005 (0.86)
Father's age	-0.002 (0.38)	-0.015* (1.69)
Mother's age	0.027*** (3.66)	0.025*** (2.89)
Father's education level	0.035*** (4.94)	0.033*** (3.88)
Mother's education level	0.037*** (5.59)	0.010 (0.91)
Urban dummy	0.242*** (4.24)	0.141* (1.88)
Observations	1920	1920
R-squared	0.89	0.85

Numbers in parentheses are absolute values of t statistics. Significance levels of 0.1, 0.05 and 0.01 are noted by *, **, and ***. All regressions include provincial dummies.

Table 11: Ordinary Least Squares Regressions Examining the Determinants of Having the Second Child in China (First-stage of 2SLS) (1989)

	Dependent variable: Two or more children dummy
Amount of fine for extra child /10000 (IV)	-0.443*** (5.81)
Proportion of girls in the household	0.012 (0.59)
Mean age of children	0.272*** (9.14)
Mean age squared	-0.012*** (8.59)
Father's age	-0.008** (2.57)
Mother's age	-0.003 (0.85)
Father's education level	-0.002 (0.47)
Mother's education level	-0.018*** (5.58)
Urban dummy	-0.090*** (3.23)
Observations	1920
R-squared	0.16

Numbers in parentheses are absolute values of t statistics. Significance levels of 0.1, 0.05 and 0.01 are noted by *, **, and ***. All regressions include provincial dummies.

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